Applied Research Branch
Strategic Policy
Human Resources Development Canada

Direction générale de la recherche appliquée
Politique stratégique
Développement des ressources humaines Canada

The Beveridge Curve and Unemployment Fluctuations in Canada

W-97-4E

by
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July 1997

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103-01758/

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First Internet Edition - 1998

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Abstract

This study applies the methodology developed by O.J. Blanchard and P. Diamond (1989) for decomposing unemployment fluctuations stemming from aggregate economic activity shocks, reallocation shocks and participation rate shocks. A structural vector autoregressive model with error correction is estimated on employment, labour force and a vacancies proxy. First of all, we demonstrate that vacancies alone react to the labour market imbalance, and that this is the only variable that introduces an error correction dynamic. Hence this variable plays an essential role in absorbing disequilibrium. The historical decomposition of structural shocks reveals the following points. The model associates strong recession-linked fluctuations in the unemployment rate with cyclical shocks. Variations in the participation rate led to an increase of more than 2 percentage units in the unemployment rate between 1972 and 1982, but this effect has dissipated completely since then. Sectorial reallocation shocks play a less significant role in unemployment fluctuations. Finally, the non-cyclical unemployment rate rose from 4% to 9% between 1970 and 1980, and has hovered around 10% since 1991. The relative contribution of cyclical and reallocation shocks is therefore comparable to the estimations made in previous Canadian studies [McCallum (1988), Burns (1990), Fortin and Araar (1997)] as well as to the conclusions of U.S. studies [Blanchard and Diamond (1989)], but estimation of the non-cyclical unemployment rate and the strong contribution of participation rate shocks are new elements.

Résumé

Cette étude applique la méthodologie développée par O.J. Blanchard et P. Diamond (1989) pour décomposer les fluctuations du chômage provenant de chocs d'activité économique agrégée, de chocs de redistribution et de chocs sur le taux d'activité. On estime un modèle vectoriel autorégressif structurel à correction d'erreurs sur l'emploi, la population active et une approximation des postes vacants. Nous montrons dans un premier temps que seuls les postes vacants réagissent au déséquilibre du marché du travail et que c'est la seule variable à introduire une dynamique de correction des erreurs. Cette variable joue donc un rôle essentiel pour résorber les déséquilibres. La décomposition historique des chocs structurels fait ressortir les points suivants. Le modèle associe aux chocs conjoncturels les fortes fluctuations du taux de chômage associées aux récessions. Les variations du taux d'activité ont induit une augmentation de plus de 2 unités de pourcentage du taux de chômage entre 1972 et 1982 mais cet effet s'est entièrement dissipé depuis. Les chocs de redistribution sectorielle jouent un rôle moins important dans les fluctuations du chômage. Finalement, le taux de chômage non cyclique serait passé de 4% à 9% entre 1970 et 1980 et à environ 10% depuis 1991. La contribution relative des chocs cycliques et de redistribution est donc semblable à ce qui à été estimé par les études antérieures canadiennes (McCallum (1988), Burns (1990), Fortin et Araar (1997)) ainsi qu'aux conclusions d'études américaines (Blanchard et Diamond (1989)) mais l'estimation du taux de chômage non-cyclique et la forte contribution des chocs de taux d'activité constituent des éléments nouveaux.

Acknowledgements

We would like to thank our two referees, Paul Storer and one anonymous person, as well as Alain Guay, Louis Grignon and David Johnson for their invaluable comments and suggestions. We also appreciate the contributions of those attending certain lunch meetings of the University of Quebec in Montreal and several members of the Applied Research Branch.

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1. Introduction

Canada's greatest macroeconomic challenge is to understand what has been keeping unemployment at such high levels since the early 1980s, and particularly since the beginning of this decade. Three main theses, not necessarily contradictory, have been developed to explain this disturbing phenomenon. The primary one argues a slowdown in economic growth, influenced by demand factors or by technological change. However, the recent coexistence of low, stable inflation with a high unemployment rate is difficult to reconcile, and is forcing an examination of some non-traditional hypotheses. For example, Fillion and Léonard (1997) point to changes in the system affecting the Phillips curve. Fortin (1996), on the other hand, argues that the presence of nominal downward wage rigidities has raised the non-accelerating inflation rate of unemployment. However, as Setterfield, Gordon and Osberg (1992) demonstrate, the Phillips curve shows amazing facility in accommodating a wide range of natural unemployment rates, opening the way for broad disagreement as to the importance of non-cyclical factors in variations in the unemployment rate.

The two main non-cyclical factors which may have pushed unemployment up are sectoral reallocation shocks and changes in the participation rate. The first empirical verifications in Canada by Samson (1985) of the Lilien thesis (1982) measured a significant contribution by sectoral reallocation shocks to unemployment fluctuations. Subsequent studies, notably by McCallum (1988) and Burns (1990), have tended, however, to conclude that this type of shock has played a small part in the largest fluctuations in the Canadian unemployment rate. Certain exogenous changes in the participation rate might have played a more important role. Moorthy (1990) and then Card and Riddell (1991) were the first to point out that, until 1990, the evolution of the employment-population ratio in Canada was sufficiently favourable to keep unemployment as low as in the United States. They suggest that the 1971 unemployment insurance reform, by stiffly increasing implicit subsidies for unstable jobs, induced a rise in the participation rate, as predicted in the classical labour supply models of Fortin (1984) and Milbourne, Purvis and Scoones (1991). The corollary to this is that the reduction of the observed participation rate since 1990 might not be solely due to the customary procyclical behaviour of the participation rate. Indeed, many changes to the unemployment insurance system since 1993 have reduced its generosity, a fact which could have had a negative impact on the participation rate.

Empirically, the difficulty lies in distinguishing between these different explanations. Abraham and Katz (1986) show that the Beveridge curve, the relationship between the unemployment and vacancy rates, can be used to separate the unemployment variations caused by aggregate activity shocks from those stemming from reallocation shocks. Whereas the former produce contrary movements of these two variables, reallocation shocks create a simultaneous increase in unemployment and vacancies. Given the strong negative short run correlation between vacancies and unemployment in the United States, Abraham and Katz (1986) conclude that aggregate activity shocks have been the predominant cause of unemployment fluctuations in that country since World War II. In Canada, Blanchard (1991) shows that the short-run fluctuations of Canadian unemployment are compatible with aggregate activity shocks. However, he points to a trend shift of the Beveridge curve which implies an increase of about 2 percentage points in the natural rate of unemployment between 1970 and 1990. Samson (1994) applies a similar analysis to five regions of Canada and obtains comparable results at both the national and regional levels. Fortin and Araar (1997) show the predominance of short-run cyclical shocks using a vector autoregression (VAR) model between the unemployment rate and a vacancies index.

Unfortunately, these analyses offer no means of distinguishing between sectoral reallocation shocks and shocks from increases in the participation rate. The present study applies the methodology developed by Blanchard and Diamond (1989) to Canadian data so that distinctions can be drawn between these three sources of unemployment fluctuations. Although analysis of the short-run joint movements of these three variables does not allow us to break down aggregate activity shocks into aggregate demand fluctuations and shocks attributable to changes in input productivity, it does provide information about the probable contribution of unemployment insurance to the unemployment rate when one takes sectoral reallocation and aggregate activity shocks into account. To this end, we estimate a structural VAR model with quarterly Canadian data from 1966 to 1996 on three aggregates, namely labour force, employment, and a vacancies proxy based on the Help-Wanted Index (HWI).

Our empirical analysis differs from that of Blanchard and Diamond (hereinafter BD) in one important regard. Despite the non-stationarity of the variables, BD still estimated a VAR in level form. This has serious consequences for estimation of the model, for as Phillips (1995) shows, the impulse responses and historical decomposition then become random and do not converge

asymptotically. When the Johansen (1988) method is used, the Canadian data support the hypothesis of a cointegration relation between employment, labour force and vacancies. We have therefore estimated a vector error correction model. In addition to ensuring convergence, this type of model can characterize long-run equilibrium and make imbalance correction mechanisms explicit.

Section 2 of this paper presents the labour market model developed by BD. Section 3 describes the data used and presents the reduced-form model. This model shows that labour market imbalance is proportional to unemployment, and that the only mechanism that works to correct the imbalance is the creation of vacancies. Without this action, the unemployment rate would not show a trend to return to a stable level. The fourth section explains how the hypotheses lead to identification of the shocks and presents the results obtained. Our findings confirm the predominance of cyclical shocks in explaining short-run fluctuations in unemployment. The historical decomposition shows that cyclical shocks could have reduced unemployment by nearly 4 percentage points in the 1970s, but a marked increase in the participation rate drove the unemployment rate up over 2 percentage points during the same period. As of 1980, however, the participation rate shocks began to reverse, with the result that in 1996 their cumulative impact on the unemployment rate was nil. As for reallocation shocks, these have had little impact on unemployment fluctuations. Finally, the model identifies a trend progression of over 4 percentage points in the unemployment rate during the period studied. The paper concludes with comments and suggestions for future research.

2. Unemployment Fluctuations and Identification of the Shocks

Here we revisit BD's theoretical model in discrete time. This model can derive identification hypotheses from the empirical model and characterize long-run equilibrium. It stands in the tradition of the models that match vacancies and unemployed persons, as initially developed by Pissarides, in which the dynamic behaviour of the labour market is generated by flows between the different labour market conditions. The extent of these flows, and certain of their properties, have been measured in Canada, notably by Jones and Riddell (1993), while Storer (1992) has demonstrated the existence of a matching function using provincial data on Alberta and Ontario.

Let labour force $L \equiv E+U$, where E is the number of jobs held and U the number of unemployed. Let the working-age population $P \equiv L+N$, where N is the number of persons not in the labour force.[1] Also defined is the total number of productive jobs $J \equiv E+V$, where V is the number of vacancies, while the number of potential jobs is $K = J+I$, where I designates unproductive jobs, i.e. vacancies which employers are not attempting to fill. In other words, of all the jobs which might theoretically exist, only those classified in J are profitable enough to be maintained or to prompt an active search for workers to fill them. The process of job destruction and creation is introduced on the assumption that, in each period, a fraction π_0 of the productive jobs become unproductive and a fraction[1] of the unproductive jobs become productive. The destruction process is assumed to affect filled jobs and vacant jobs at the same rate, and it is further assumed that the jobs being created are initially vacant and will be filled by a process of matching unemployed persons with vacancies.[2] The equation of motion of J can therefore be written:

$$\Delta J = \pi_1 I - \pi_0 J = \pi_1 (K - J) - \pi_0 J = \pi_1 K - (\pi_0 + \pi_1) J$$

[1] For purposes of simplicity, we situate our model in a stationary economy. Adaptation to a growth economy is direct if all stock variables are interpreted as per capita variables. For example, dividing by P, L and N would be interpreted as participation and non-participation rates respectively, while E would be the employment-population ratio. The model of a growth economy would contain an additional term to capture the trend growth of the population.

[2] The model could be modified to include non-null transitions from I to E. Since this would weigh down the model without fundamentally changing its properties, it seemed to us preferable to maintain the hypothesis as formulated.

The steady state is then characterized by: $\frac{J}{K} = \frac{\pi_1}{\pi_0 + \pi_1}$

Let $c = \pi_1/(\pi_0 + \pi_1)$, the relation that exists at steady state between productive jobs and the potential number of jobs. We interpret c as a measure of the cyclical situation. However, our use of the word cyclical must be defined. The theoretical framework does not allow for a distinction between aggregate demand shocks and shocks to productivity. We use the word cyclical to designate any fluctuation in aggregate activity which translates into a larger proportion of profitable jobs in the economy, whether these fluctuations stem from aggregate demand or from shocks to productivity.

At steady state, total flows from J to I and vice versa are $\pi_0 J = \pi_0 cK$. Dividing these flows by K gives the intensity of the flow of sectoral job reallocation. The intensity of this flow serves to define the indicator of sectoral shocks $s = \pi_0 c$. Variables s and c are therefore two of the three sources of shocks that we want to distinguish empirically. The rest of the model will serve to derive the impact of these shocks on variables E, U and V. Note that a change in the transition rates produces movements in c and s which are not independent. Hence a permanent increase of π_0 reduces c and increases s at steady state. To obtain an increase of s without changes in c, π_1 has to increase at the same time as π_0 so as to keep the $\pi_1/(\pi_0 + \pi_1)$ ratio constant.

When filled jobs become unproductive, the employer-employee relation is broken and the employee becomes unemployed.[3] The job destruction process thus leads to a flow of $\pi_0 E$ in every period. We also assume that transitions from E to U occur due to quits at a rate q. Flows to E occur

[3] This model implies that direct transitions between E and N are null, and that there would be a necessary movement through U when such a transition occurs. As Jones and Riddell (1993) demonstrate, this transition rate is not null, since nearly 40% of those out of the labour force who are moving into the labour market or vice versa are not measured as unemployed persons in the meantime. Taking these additional transitions into account considerably complicates the model's algebra without fundamentally changing the behaviours we want to highlight. For this reason we will maintain this hypothesis in the presentation of the theoretical model. However, these transitions will be taken into account in the empirical analysis.

solely because of a process of matching unemployed persons with vacancies. This process is described by a function $\alpha m(U,V)$, where α is an indicator of the fluidity of the process. It is assumed that $m_U>0$, $m_V>0$, and $U \cdot m(V/U) = m(U,V)$, i.e. that the matching function displays constant returns to scale.[4]

Finally, we assume that movements of those out of the labour force into the labour market translate a procyclical behaviour and exogenous shocks f to the participation rate. The behaviour of the participation rate can thus be written $\Delta L = a\Delta E + f$, with $0 < a < 1$. The equations of motion of variables E, U and V can thus be written:

$$\Delta E = \alpha m(U,V) - (q + \pi_0)E \tag{1}$$

$$\Delta V = -\alpha m(U,V) + \pi_1(K - V - E) + qE - \pi_0 V \tag{2}$$

$$\Delta U = (a-1)\Delta E + f = -(1-a)\alpha m(U,V) + (1-a)(q+\pi_0)E + f \tag{3}$$

These three equations describe the system's behaviour when it is disturbed by c, s and f shocks. To render the dynamics of these shocks, parameters π_0 and π_1 are replaced by s/c and $s/(1-c)$ respectively,[5] which gives:

$$\Delta E = \alpha m(U,V) - (q + s/c)E \tag{4}$$

$$\Delta V = -\alpha m(U,V) + [s/(1-c)](K - V - E) + qE - (s/c)V \tag{5}$$

$$\Delta U = -(1-a)\alpha m(U,V) + (1-a)(q+s/c)E + f \tag{6}$$

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[4] The hypothesis of constant returns to scale is necessary in order to be able to interpret the stationary model as a rewriting of a balanced growth model.

[5] Knowing that $c = \pi_1/(\pi_0+\pi_1)$, this implies that $\pi_1 = c\pi_0/(1-c)$. The expression indicated is then obtained by substituting s for $c\pi_0$.

As BD show, this model always possess a stable dynamic.[6] Steady state is obtained by requiring that $\Delta E = \Delta V = \Delta U = 0$, which gives the system of equations:

$$am(U,V) = (q + s/c)E \tag{7}$$

$$am(U,V) = [s/(1-c)](K - V - E) + qE - (s/c)V \tag{8}$$

$$am(U,V) = (q + s/c)E + f/(1-a) \tag{9}$$

Substituting equation 7 in equation 9 implies that $f = 0$, i.e. that for a steady state to exist, shocks to the participation rate have to be null. In other words, the existence of a steady state depends entirely on the existence of a stable long-run participation rate, something which is guaranteed in the long term.[7] Equations 7 and 8 make it possible to recover the equation $J/K = c$, which is the equilibrium between productive jobs and potential jobs. These two equations also imply that variables U, V and E are linked by a relationship that generates long-run equilibrium. Consequently, if it can be demonstrated that these variables are integrated of order 1, this set of equations implies the presence of a cointegration relation.[8] This is an important property for estimating the model, and will give rise to an empirical verification.

The short-run dynamic of adjustment to the different types of shocks is intuitively easy to define. After a shock to c, the new steady state is at a lower level of U and a higher level of V and E. The more favourable situation is initially manifested through a higher level of vacancies. Since the growth of V increases the number of matches, the result is a subsequent rise of E and a parallel reduction of U. The decline of U is, however, smaller than the rise of E since, as employment increases, more people enter the labour market. Furthermore, since the movements of V precede

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[6] The stability condition is that the slope of $dV/dt = 0$ in plane U V is greater than that of $dU/dt = 0$. By taking the total differential relative to U and V in equations 8 and 9 and comparing the slopes of these two curves, one can show that the stability condition is still satisfied.

[7] Since the participation rate is a positive fraction less than one, it is impossible for that rate to be continually increasing or decreasing. In practice, as shown by the evolution of the female participation rate, it may take more than one generation for the participation rate to stabilize. Therefore steady state may not be observed in a finite sample.

[8] Random variables are said to be cointegrated of order k, noted I(k), if it is possible to find k different combinations of random integrated variables with an order of integration less than that of each variable taken separately. If variables U, V and E are I(1), combination $m(U,V) - (q+s/c)E = 0$ has this property since 0 is integrated of order zero.

those of U, the model calls for a cycle which is visually characterized by a counterclockwise loop in the Beveridge space.

A sectoral job reallocation shock displaces the system to a new steady state characterized by a higher level of U and V along a 45° straight line. In addition, since J is presumed constant, the short-run dynamic implies that on impact, one vacancy appears for each job lost. Since the theoretical model does not permit direct transitions between E and N, all employees who lose their jobs become unemployed, and hence there should be a simultaneous, identical rise in U and V. In practice, however, as demonstrated by Jones and Riddell (1993), direct flows from E to N do occur, so that sectoral shocks raise both unemployment and the number of persons not in the labour force. Empirically, therefore, we must expect the increase in unemployment that accompanies these shocks to be less than the increase in vacancies. In other words, a sectoral shock translates empirically into a shift with a slope of less than 45% in plane U, V, which is consistent with what BD have found for the United States.

We close our analysis of the different types of shocks by describing the effect of shocks to f. On impact, these produce a horizontal displacement in plane U, V. Furthermore, since curve $dU/dt=0$ shifts right, the new steady state is characterized by a higher level of U and a lower level of V.[9] These shocks empirically correspond to changes in L that are not related to changes in E. Given the trend increase in the female participation rate, the effect of which has been only partially offset by the trend decrease in the male, the Canadian data will probably reveal a trend shift in the equilibrium relation between U and V.

In summary, the theoretical model is a means of indicating the system's dynamic responses to the different types of shocks. It also shows that vacancies play a key role in correcting labour market disequilibrium. We use the variables' projected reactions to the various shocks as our basis for empirical identification of the shocks. The next section describes the variables as well as the empirical model we are estimating.

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[9] Remember that the slope of $dV/dt = 0$ in plane U,V is greater than the slope of $dU/dt = 0$. Since both are negatively sloped, $dU/dt=0$ is more horizontal than dV/dt and the intersection at the new steady state is thus below and to the right of the old one.

3. The Data and the Reduced-Form Model

We estimate the model with seasonally adjusted data on employment, labour force and vacancies. The *E* and *L* series are from the Labour Force Survey. These data were adjusted by Statistics Canada in 1994 and the CANSIM labour force series began in 1976. To obtain continuous series since 1966, we have consolidated these series with those published in CANSIM prior to the 1994 adjustment.[10] However, there is no current series on the number of vacancies in Canada. A vacancies survey was carried out by Statistics Canada between 1971 and 1978 but no national estimate has since been done.[11] There is the same difficulty in the United States. BD have used the estimation of vacancies produced by Abraham (1987) based on the HWI. This estimation has shown that the HWI is a good indicator of short-run vacancy movements.

The HWI has been estimated in two different ways in Canada. From 1962 to 1988, it was proportional to the space occupied by job offers in the country's leading daily newspapers. The current series, available since 1981, instead indicates the number of job offers in those newspapers, the same concept as is used in the United States.[12] To obtain a continuous series over a 30-year period, we consolidated these two series by effecting a reverse flow forecast of the number of jobs advertised prior to 1981, the details of which are described in Appendix 1.[13] After obtaining a continuous HWI series for 1966-1996, we created an estimate of the number of vacancies by adjusting the series level so that it was identical to the mean number of vacancies estimated by Statistics Canada between 1971 and 1978, namely 127,600.

Based on the data on *L*, *E* and *V* (see figure), we can calculate the unemployment rate $UR=100*(L-E)/L$ and the vacancy rate $VR=100*V/(E+V)$. From these rates we trace the Beveridge

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[10] If *LEN* is the logarithm of the new employment series and *LEO* is the former series, values prior to 1976 have been adjusted by the equation *LEN* = -0.54861508 + 1.0632822**LEO* which has been estimated on the monthly data from January 1976 to April 1994. The labour force logarithm has been adjusted by the equation *LLN* = -0.53398816 + 1.0611226**LLO*.

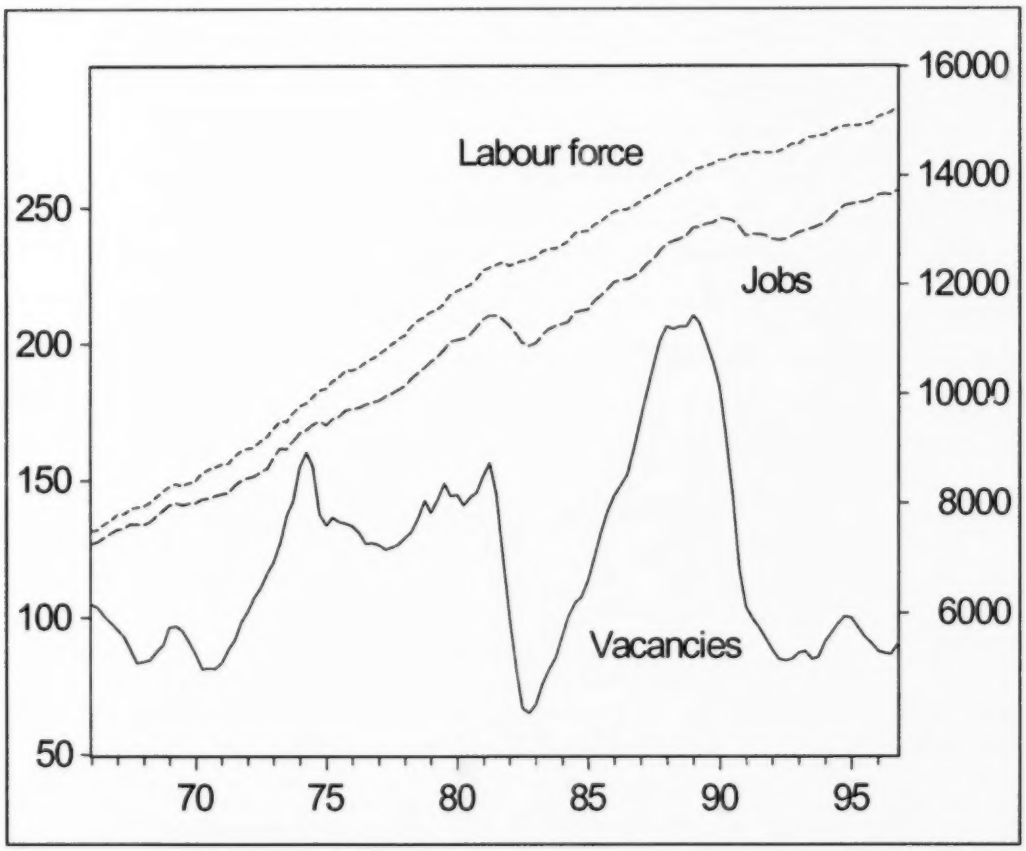
[11] According to a 1995 Statistics Canada survey in Quebec to estimate employers' hiring intentions, the estimated number of vacancies was approximately 31,000. This would correspond to a vacancy rate of less than 1%.

[12] The current series in Canada is thus conceptually identical to that used in the United States.

[13] Like Zagorsky (1993), we proxy vacancies in Canada using a consolidated HWI series. Our consolidation is different however, since Zagorsky uses a forecast from the old series to obtain a continuous annual series from 1962 to 1990. His

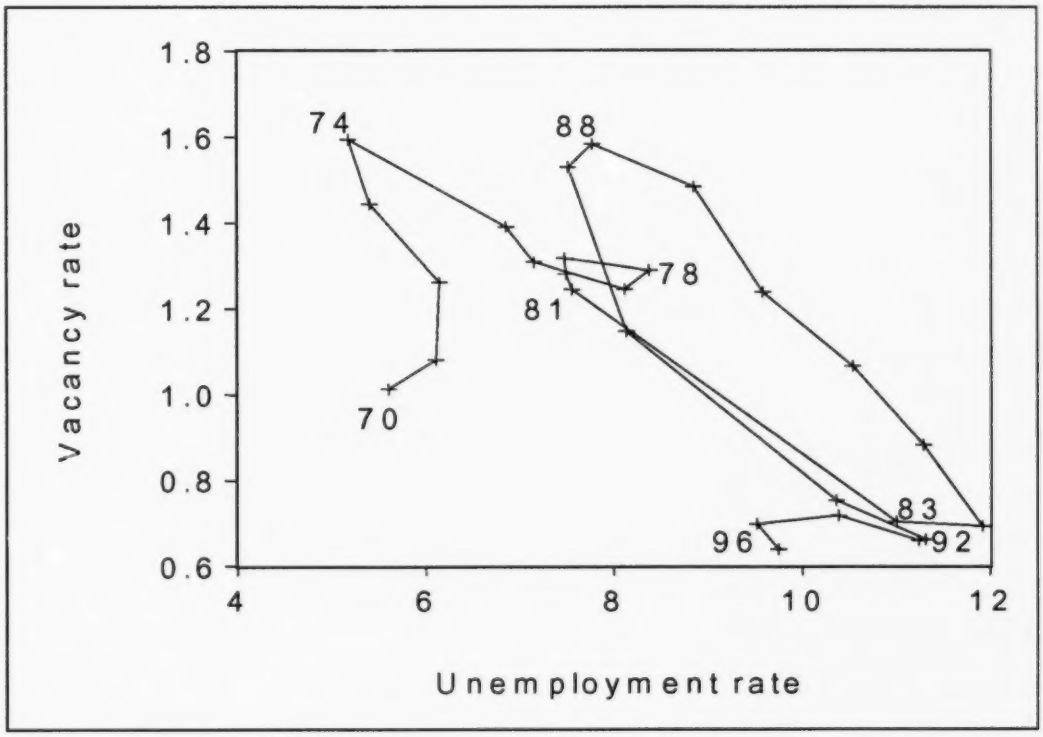
curve presented in Figure 2, which reveals two major facts. First, the biggest variations in the unemployment rate are associated with a shift along a counterclockwise loop between 1982 and 1992, typical of a complete economic cycle. Second, the Beveridge curve slides to the right between 1970 and 1976, but this movement is reversed starting in 1994. Since such lateral shifts are not consistent with the foreseeable consequences of aggregate activity shocks, they suggest that reallocation shocks or participation rate shocks caused unemployment to increase in the early 1970s and decrease starting in 1994. The empirical analysis aims to clarify the respective contribution of these shocks.

Figure 1 The Data (000)



choice was motivated by the fact that he had to forecast only two years to obtain a continuous series. We preferred a method that allows for the addition of new observations as they become available.

Figure 2 The Beveridge Curve (1970-96)



Although the 1966 to 1996 series are available monthly, we decided to estimate the model on quarterly data. This gives a more parsimonious representation without significantly reducing the power of the tests of hypotheses, since it depends more on the length of the sampling period than on frequency. The sample contains 124 quarterly observations starting in 1966:1 and ending in 1996:4. We chose to estimate the model using the variables *L*, *E* and *V* rather than variables *U*, *E* and *V*, since the model then has more desirable statistical properties.[14]

[14] With level estimation, choosing between the two sets of variables would be of no consequence, since $U = L - E$. However, a logarithmic transformation is required to make the series' variance stationary, and the result of this non-linear transformation is that the two representations are no longer equivalent. When estimating a model in *u e v*, the *u* residuals present a distribution that is too flat. By comparison, models estimated with *l e v* pass the test of normality with no problem. We attribute the excessive flattening of the distribution of unemployment to the logarithmic transformation.

To determine the model type appropriate to the data, we first verified the order of integration of the variables. Table 1 presents the results of the augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests done on the logarithm of the variables.

Table 1 Unit Root Test

Variable	ADF (level)	ADF (diff)	PP	Critical Value of 1%[a]	Critical Value of 5%[a]
log(*L*)	0.3880	-10.9676	0.2650	-4.0355	-3.4469
log(*E*)	-2.1172	-5.2208	-1.9299	-3.4852	-2.8853
log(*V*)	-2.6671	-4.5785	-2.2995	-4.0367	-3.4475

[a] MacKinnon's critical values for rejection of null hypothesis of a unit root.

Note: The tests were done using one lag for *e*, 2 for *v* and none for *l*, so as to eliminate autocorrelation of the residuals. To determine whether or not to adopt the deterministic trend, we initially conducted the test with one constant and one deterministic trend. For the 3 variables, the null hypothesis that the root is unitary could not be rejected. We then tested the joint hypotheses of a unit root and a deterministic trend. For variables *l* and *v*, this joint hypothesis was rejected, whereas it could not be rejected for *e*. The equation thus contains one deterministic trend and one constant for *l* and *v* but only one constant in the *e* equation.

The null hypothesis of a unit root is not rejected for each series, but this hypothesis is clearly rejected in the first difference. Since these variables are I(1), the level estimation can create spurious correlations. We therefore checked whether these variables are cointegrated using Johansen's procedure (1988), which requires that a vector error correction model be estimated in the form:

$$\Delta x_t = \mu_0 + \Sigma_{i=1}^{p} \Gamma_i \Delta x_{t-i} + \prod x_{t-1} + \mu_t \qquad (10)$$

In this equation, $x = (l\ e\ v)'$ is a 3×1 vector of observations at time t, μ_0 is a 3×1 vector of constant terms, Γ_i is a series of 3×3 matrices p capturing the short-run dynamic at lag i, Π is a 3×3 matrix capturing the long-run effect, and u_t is a 3×1 vector of error terms ($u_l\ u_e\ u_v$). The Johansen procedure decomposes the Π matrix into a product of two matrices ($\alpha\beta'$) of dimension 3×r, where r is the number of cointegration relations. Matrix β is comprised of the r cointegration vectors, and product $\beta' x_{t-1}$ is the stationary subspace of dimension r formed by these cointegration vectors. Element ij of matrix α indicates the importance of the j^{th} cointegration relation in the i^{th} equation.

This is a measure of the speed at which short-run disturbances are absorbed and the system returns to long-run equilibrium. The number of cointegration relations is determined by a test on the rank of matrix Π. This rank gives the dimension of the stationary subspace and is therefore identical to the number of cointegration relations.

Different formulations of the model's deterministic components are possible. Models which have no deterministic components or which permit quadratic trends were excluded a priori because they are not very plausible. We therefore compared a general model with a constant in the first difference of the series and a trend in the cointegration relation, a model with no deterministic components in the cointegration relation but with a constant in the difference of the variables, and finally a model with a constant in the cointegration relation but no deterministic elements in the first difference of the series. Systematic comparison of the different models by the Pantula principle is a means of sequentially testing the non-rejection of a general model versus a more restrictive model. The two types of tests proposed are the λ_{max} test and the trace test.[15]

Although crucial for the rest of the empirical work, this selection stage is difficult. The tests for selecting the model are very sensitive to p, the number of time lags necessary to capture the short-run dynamic. We include in the model a sufficient number of lags for the residuals to be normally and independently distributed. To verify these properties, we therefore varied the number of lags between 2 and 12 quarters and analysed the statistical properties of the residuals by using autocorrelograms up to 24 lags and by applying the LM test at 1 and 4 lags, the Ljung-Box test at 24 lags and the Bera and Jarque test for normality. However, since the choice of the model itself affects the conclusions of the autocorrelation and normality tests, the conclusions of these tests have to be compared according to the different possible types of models. In short, a complete grid exploration of the various combinations of models and lags has to be carried out.

[15] The trace test is based on the statistic $\lambda_{tr} = -T\Sigma\ln(1-\lambda_i^*)$, where λ_i^* are the estimated values of the characteristic roots of matrix Π.

This grid analysis shows that the model type has very little influence on the number of lags suggested, since in all cases 6 lags were sufficient for the residuals to be normally and independently distributed. Table 2 reports the results of the model specification tests, of which the only deterministic elements are the constants in the first differences of the series. Clearly, the normality of the residuals is never rejected, whatever the number of lags. On the other hand, when the number of lags is below 6, significant residual autocorrelations persist. These reappear when the number of lags reaches 10 or more. When it varies between 6 and 8, one cannot reject the hypotheses of independence and normality of the residuals. Since the representation with 6 lags is more sparing, this is the one we adopted.

Table 2 Analysis of Residuals

Type of Test	$p = 2$	$p = 4$	$p = 6$	$p = 8$	$p = 10$	$p = 12$
L-B (p-val)	0.06	0.04	0.12	0.16	0.04	0.00
LM(1) (p-val)	0.98	0.07	0.12	0.76	0.04	0.53
LM(4) (p-val)	0.06	0.05	0.68	0.08	0.00	0.48
Normality (p-val)	0.21	0.38	0.65	0.30	0.34	0.38

Rank tests applied to the model with 6 short-run matrices lead to the selection of a model with a cointegration relation whose only deterministic elements are the constants in the variables' growth rate equations. Table 3 shows how the rank tests vary according to the number of lags. Generally speaking, when there are less than 6 lags the tests detect 2 cointegration relations, and just one relation when there are 8 or more lags. When 6 lags are included, the conclusions are not so clear. The trace test, calculated with a single cointegration relation, is 11.87 even though the critical value at 90% is 13.31, which does not allow for the rejection of this model. On the other hand, λ_{max} calculated from the same model is greater than the critical value of 10.60, suggesting the presence of two cointegration relations. Despite everything, we conclude that a single cointegration relation is present. Two of the system's characteristic roots have a value of 0.9981, i.e. on the unit circle, while the other roots have a norm below unity. The system is thus subject to two joint stochastic trends, as is compatible with the presence of a single cointegration relation.

Table 3 Rank Test

	Lambda max test						Critical values
	$p = 2$	$p = 4$	$p = 6$	$p = 8$	$p = 10$	$p = 12$	At 90 %
r = 0	35.5	28.7	19.9	19.8	16.7	20.5	13.4
r = 1	17.9	23.8	11.8	7.4	8.1	4.3	10.6
r = 2	0.2	0.5	0.1	0.1	0.2	0.0	2.7
	Trace test						Critical values
	$p = 2$	$p = 4$	$p = 6$	$p = 8$	$p = 10$	$p = 12$	At 90 %
r = 0	53.7	52.5	31.8	27.2	24.9	24.8	26.7
r = 1	18.2	24.4	11.9	7.4	8.2	4.3	13.3
r = 2	0.2	0.5	0.1	0.1	0.2	0.0	2.7

The constrained model in which a cointegration relation is imposed is presented in Appendix 2. Imposition of the cointegration relation does not degrade the statistical properties of the model. In particular, the residuals are still normally distributed white noise, and in requiring the model to have 2 unit roots, the other roots remain inside the unit circle. This model was then estimated recursively from 1975:4. The recursive analysis provided additional information on the number of cointegration relations and the stability of the parameters. The recursively calculated trace test shows only one cointegration relation increasing with time, as is consistent with the existence of a single cointegration relation. The log likelihood, remaining inside the confidence interval, does not demonstrate any system instability problems. The one ahead period prediction test applied to the individual series detects an instability in the forecasting of vacancies in 1981, with 4 predictions outside the confidence interval. However, the ratio of estimated beta to known beta recursively confirms that the relation is not subject to instability problems. On the whole, then, the problems with the specification chosen seem to be minor.

We first look at the degree of dependence between the variables and the shocks of the reduced-form model. Table 4 gives the covariances of d(e), d(v), d(I), the covariances of the shocks, and the correlations, in italics, in the upper triangular part of the tables. There is a strong correlation

of 0.74 between the shocks to e and l, symptomatic of the procyclical behaviour of the participation rate. However, the correlation between u_v and u_e is only 0.31, and that between u_v and u_l is practically zero.

Table 4 Covariances and Correlations of the Reduced-Form

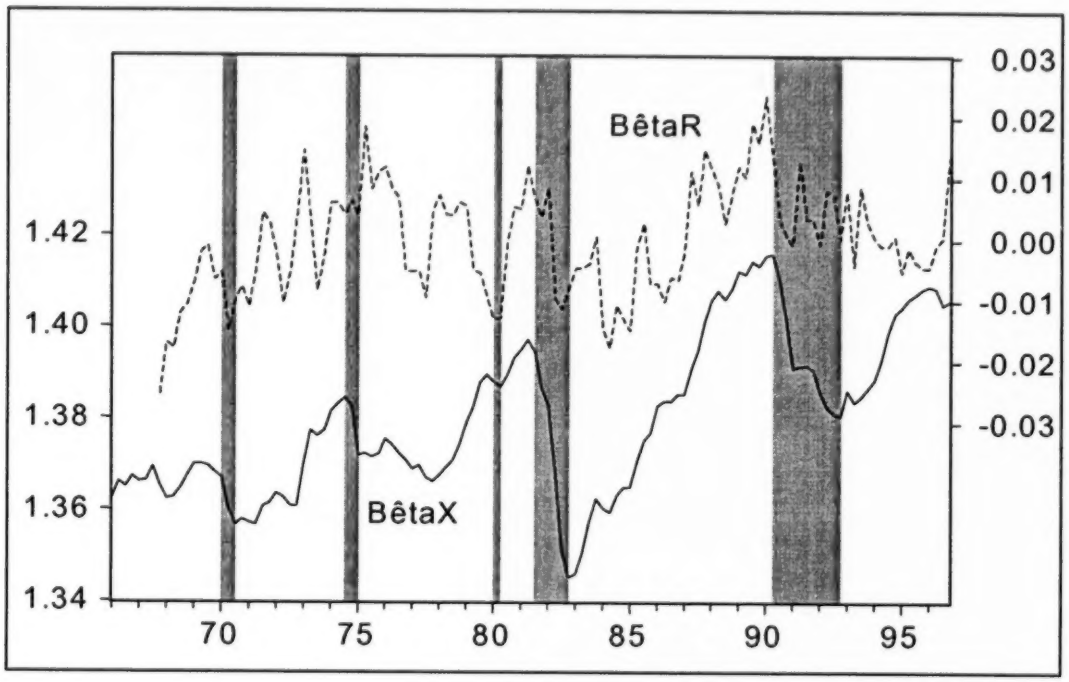
	d(*e*)	d(*e*)	d(*e*)
d(*e*)	0.0000404	0.6617	0.7417
d(*v*)	0.0002512	0.0003566	0.2446
d(*l*)	0.0000211	0.0000653	0.0000200
	u_e	u_v	u_l
u_e	0.0000135	0.3058	0.7423
u_v	0.0000365	0.0010538	0.0094
u_l	0.0000088	0.0000010	0.0000105

The cointegration relation is of the form $\beta' X = \beta_e e + \beta_v v + \beta_l l$. After normalization relative to e, this relation is written e - 0.009×v - 0.839×l. This equation provides a gross measure of labour market imbalance, "gross" because it recovers the impact of the deterministic variables as well as the transitory effects related to the model's short-run dynamics. When these impacts are removed from it, one has a net measurement of the imbalance, purged of these transitory and deterministic effects. These two measurements are presented in Figure 3.[16] The grey areas indicate the recession dates identified by Statistics Canada according to per capita GDP.[17] Note that the gross measurement is strongly correlated with the economic cycle, whereas this correlation is much less visible with the net measurement. According to the net measurement, the largest negative

[16] If R defines the vector formed of the residuals of the regressions by ordinary least squares of e, v and l on the deterministic variables and the variables lagged in difference, the net measurement of the imbalance is equal to $\beta' R$.. It is this latter series, containing much less persistence, that is tested for stationarity in the Johansen procedure.

disequilibrium were observed in 1979 and 1983, while the largest positive disequilibrium were in early 1975 and late 1989.

Figure 3 Measurements of Labour Market Imbalance



To better understand the system's adjustment dynamic, we explored how the cointegration equation works by conducting a series of tests on β and α, presented in Table 5. Test LR shows that the nullity of the impact of *e* and *l* in the cointegration relation is rejected at marginal levels 0.01 and 0.02 respectively. However, the marginal level of the same test is only 0.50 for *v*, which implies that this variable has no significant impact in the cointegration relation. Therefore, even if vacancies theoretically play a role in the long-run relation, the impact of this variable is not statistically

[17] See the February 1996 issue of the *Canadian Economic Observer,* which compares different criteria for dating the recessions in Canada.

significant. In economic terms, this means that labour market imbalance is related mainly to the difference between *e* and *l*, a value strongly tied to the unemployment rate.

Table 5 Tests on β and α

Null Hypothesis (H1 : ≠ 0)	Type of Test	Value Calculated	Marg. Probability
Test H0: $\beta_e = 0$	Chi2(1)	6.420	0.01
Test H0: $\beta_v = 0$	Chi2(1)	0.450	0.50
Test H0: $\beta_l = 0$	Chi2(1)	5.820	0.02
Test H0: $\alpha_e = 0$	Student's t	0.863	0.39
Test H0: $\alpha_v = 0$	Student's t	-2.568	0.01
Test H0: $\alpha_l = 0$	Student's t	-0.455	0.65
Test H0 : $\beta_v = \alpha_e = \alpha_e = 0$	Chi2(3)	2.560	0.46

We then verified another important element, namely how the variables react to the imbalance, by testing the nullity of the coefficients of α. Considered individually, one cannot reject the null hypothesis that variables *e* and *l* do not react to the imbalance, since the marginal levels of the Student's *t*-tests are only 0.39 and 0.65 respectively. On the other hand, the null hypothesis that $\alpha_v = 0$ has a marginal significance level of 0.01. Hence this is the only variable whose speed of adjustment to equilibrium is not statistically null. In other words, only vacancies react directly to the imbalance, whereas employment and labour force are weakly exogenous relative to the cointegration vector. The individual tests thus suggest that α_e, α_l and α_v are null. This conclusion is confirmed by the simultaneous nullity test of these three parameters whose marginal level is 0.46. Table 5 presents the F's calculated and the marginal significance levels of the simultaneous nullity tests of the lags of d(*e*), d(*v*) and d(*l*) in the three equations. These tests show that the d(*v*) lags are very significant in the three equations but the d(*e*) and d(*l*) lags have no statistically significant effect.

The economic interpretation of these various tests is particularly interesting, and highlights the key role of vacancies and matching in the correction of labour market disequilibrium. First of

all, the fact that v is not significant in the cointegration relation means that only the difference between e and l, a value directly related to the unemployment rate, is symptomatic of the imbalance. However, since these two variables do not react directly to the imbalance, unemployment does not tend to return to a stable level by itself. Only the creation of vacancies serves to correct the labour market imbalance. In fact, only v is sensitive to the imbalance between e and l, and it is also the only variable that introduces a short-run dynamic into the system. The creation of vacancies thus leads to subsequent adjustments in e and l which help to gradually eliminate the imbalance, that is, unemployment. These results therefore provide important support for the job search and matching model, for without this mechanism, the unemployment rate would be characterized by hysteresis.

4. The Structural-Form Model

The reduced form makes it possible to estimate the interaction between the system variables, to measure the labour market imbalance, and to test the imbalance correction mechanisms. However, it does not provide information about the origin of the disturbances. We now turn to the theoretical restrictions presented in section 2 as a basis for identifying the sectoral job reallocation shocks *s*, aggregate activity shocks *c* and labour force shocks *f*. Let the vector of structural shocks $\varepsilon = (s\ c\ f)'$ which are not serially correlated.[18] The contemporaneous impact of the structural shocks on the reduced-form shocks $u = (u_e\ u_v\ u_l)$ comprises the elements of matrix A such that $u = A\varepsilon$. For example, the matrix of the variances of reduced-form shocks $\Sigma = E\{uu'\}$ and that of structural shocks $\Omega = E\{\varepsilon\varepsilon'\}$ are related by the equation $\Sigma = E\{A\varepsilon\varepsilon'A'\} = A\Omega A'$. The model can be identified by a set of restrictions on A or on ε. Following the restrictions described in section 2, we define A in this form:

$$\begin{bmatrix} \mu_e \\ \mu_v \\ \mu_l \end{bmatrix} = \begin{bmatrix} -1 & \alpha & \lambda \\ \beta & 1 & 0 \\ -\theta & \theta\alpha & 1 \end{bmatrix} \begin{bmatrix} s \\ c \\ f \end{bmatrix} \tag{11}$$

The elements on the main diagonal are simply normalizations. They indicate that at impact, the effect on u_e of a shock to *s* is unitary, that a shock to *c* increases u_v by 1, and a shock to *f* increases u_l by 1. The coefficient θ is a fraction that reflects the systematic reaction of the labour force to employment changes, while λ captures the impact of the direct flows between non-participation and employment. Coefficients α and β characterize the cyclical shocks and the sectoral reallocation shocks. The value given to α indicates the effect of a cyclical shock on u_e whereas β shows how much u_v increases following a shock to *s*. Finally, in conformity to the

[18] A formal distinction should be made between structural disturbances and the shocks to these disturbances purged of the effects of their own dynamic. As BD point out, however, they cannot be empirically distinguished. The only dynamic that can be observed is the one captured by the reduced form, which is a convolution of the structural disturbances dynamic and the system dynamic. The shortcut we take here is a valid representation on the hypothesis that the structural disturbances are white noise processes.

theoretical model, it is assumed that shocks to *f* have no contemporaneous impact on *v*. The whole identification process will take place by combining different values of the parameters α, β, λ and θ.

Two main approaches are possible to identify the model. Like King, Plosser, Stock and Watson (1991), we could use the fact that, since the model has a cointegration relation, the system is subject to one shock which has only temporary impacts and two shocks which have permanent impacts. Restrictions on the long-run variance matrix would then make it possible to decompose the shocks. In such an approach, it would be natural to associate the temporary shocks with cyclical shocks. However, this would not be compatible with our initial interpretation of cyclical shocks, which cover shocks traditionally associated with aggregate demand, particularly monetary shocks whose impacts are temporary, as well as technology shocks whose effects are likely to be permanent in part. We have therefore adopted the approach used by BD, namely to identify the shocks on the basis of impulse responses.

The solution to the identification problem is partially sequential. We must first find solutions for the values of θ and λ. Coefficient θ can be estimated by regressing u_l on u_e, producing an estimate $\gamma = 0.6519$ of the procyclical reaction of the participation rate. However, γ is a biased estimate of θ when $\lambda \neq 0$ because it also captures the effect on employment of shocks to *f*. A non-linear equation can correct this bias if we postulate a value for λ, since the coefficients θ and λ constitute a pair of variables determined together, but separately from the rest of the system of equations.[19]

We have little information on the relative extent of direct flows between non-participation and employment following shocks to *f*. Jones and Riddell (1993) have estimated these flow rates, as well as the flow rates between unemployment and non-participation. Their results suggest that the total flow between *N* and *E* is about half the flow between *N* and *L*.[20] While these values are

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[19] This equation is $[\gamma - \theta](1 - \theta^2 - \lambda^2) = [(\sigma_l^2/\sigma_e^2)\theta^2]\lambda(1 - \theta\lambda)$. It is one of the equations obtained in solving the system $\Sigma = A\Omega A'$.

[20] We have used the March 1984 data from their sample, which provides details on the size of the different groups. By weighting the flow rates of marginally attached and unattached persons, one finds that the probability of flow during one month from *N* to *E* is 0.0323, while the probability from *N* to *U*+*E* is 0.0657, for a ratio of 49.1%.

suggestive of the relative extent of the direct transitions between N and E, they are based on observed flow rates and not those resulting solely from shocks to f. Fortin and Fortin (1997) have estimated determinants of the participation rate for different demographic groups in five regions of Canada. They correct the problems inherent in the simultaneity bias by using an instrument to estimate the impact of the employment rate on the participation rate. They demonstrate that the participation rate's sensitivity to change in the employment rate varies between 0.20 and 0.70, depending on the different groups. For purposes of identification, we therefore adopt a mean value of $\lambda = 0.4$, which implies a value of $\theta = 0.4623$.[21]

Once the values of λ and θ have been determined, coefficients α and β are chosen so that the responses of the system variables to s and c shocks are consistent with theoretical expectations over as many quarters as possible. Pairs of values are sought which produce a response to the s shocks that is positive for v and negative for e, although the response of these two variables to a c shock must remain positive. The values adopted for identification are $\alpha = 0.0544$ and $\beta = 2$. These were adopted through a grid exploration of the different parameter combinations, given the values imposed on θ and λ. This exploration of the different possibilities revealed that the values of α and β that produce impulse responses to an s shock compatible with theoretical expectations lie in a very narrow band. When one deviates from the values indicated, the responses of the two variables to reallocation shocks are generally both positive, and hence similar to the c shocks. On the other hand, it is also very difficult to find values that produce a response to a c shock that is different from theoretical expectations. In other words, the combination chosen is the most plausible scenario for the simultaneous presence of aggregate activity shocks and cyclical shocks.[22]

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[21] BD use a value of 0.50 instead.

[22] BD have also noted the difficulty of finding a pair of values capable of producing impulse responses consistent with theoretical expectations of reallocation shocks. The values they find are very different from those we have used. However, we later outline the responses to shocks in the Beveridge space. These responses are very similar on impact. The difference in coefficients between the two studies probably arises from the data transformations, since BD normalize their observations and reduce the variance of the three series so that it is unitary.

Our identification hypotheses differ from BD on one important point. Unlike them, we do not assume that the structural shocks are orthogonal, for two reasons. First, if one assumes that Ω is a diagonal matrix, this implies a θ value of only 0.11,[23] which seems to us much too low when one considers the information from other studies. This means that the behaviour of the reduced form is not compatible with the hypothesis of orthogonal structural shocks. Apart from this primary reason, the hypothesis of orthogonality among the structural shocks seems to be invalidated by the work of Davis and Haltiwanger (1991). Using longitudinal data, they show that job displacements are negatively correlated with cyclical fluctuations. In terms of the theoretical model in section 2, this means that changes to flow rate π_0 are not accompanied by changes of π_1 such that $\pi_1 /(\pi_0+\pi_1)$ remains constant. In order to obtain impulse responses from the structural-form system all the same, we therefore orthogonalized the structural shocks by a Choleski decomposition, imposing the order *s, c, f*. Given the difficulty of finding parameter values compatible with the theoretical effects of sectoral shocks, we placed the reallocation shocks first to give the greatest possible contribution to the sectoral shocks. Shocks to *f* have been placed last, so as to measure the minimum contribution that can be assigned to labour force shocks in unemployment fluctuations.

Table 6 gives the covariances and correlations of the structural-form shocks. Despite the fact that the shocks were not constrained to be orthogonal, the correlations of the structural shocks remain weak. For example, the strongest correlation appears between u_f and u_c and is only –0.271, that between u_f and u_s is –0.156, and that between the sectoral and cyclical shocks is only –0.034. Because of these low correlations, the order in which one performs the Choleski decomposition will have but a limited impact on the estimated contribution of each shock.

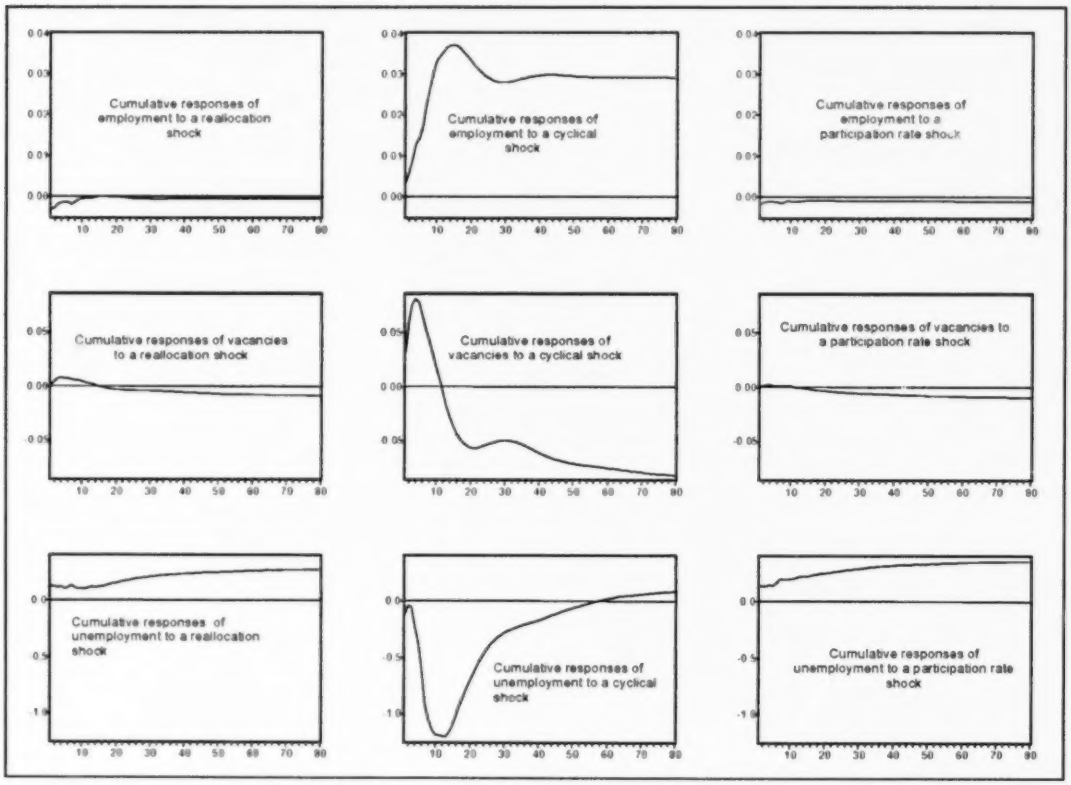
[23] When one solves the system $\Sigma = A\Omega A'$ on the hypothesis that the shocks are independent, one obtains a system which is in part recursive and where the first equation to solve is $\text{cov}(u,u_l) = \theta\text{cov}(u_c u_l)$, implying that $\theta = 0.000001/0.0000088$. Once this value of θ is found, one solves for λ. Curiously, BD have not solved for θ in this way, but instead used an auxiliary regression, as we did.

Table 6 Covariances and Correlations of Structural Shocks

	u_s	u_c	u_f
u_s	0.00000974	-0.03420000	-0.15570000
u_c	-0.00000338	0.00100670	-0.27110000
u_f	-0.00000120	-0.00002129	0.00000613

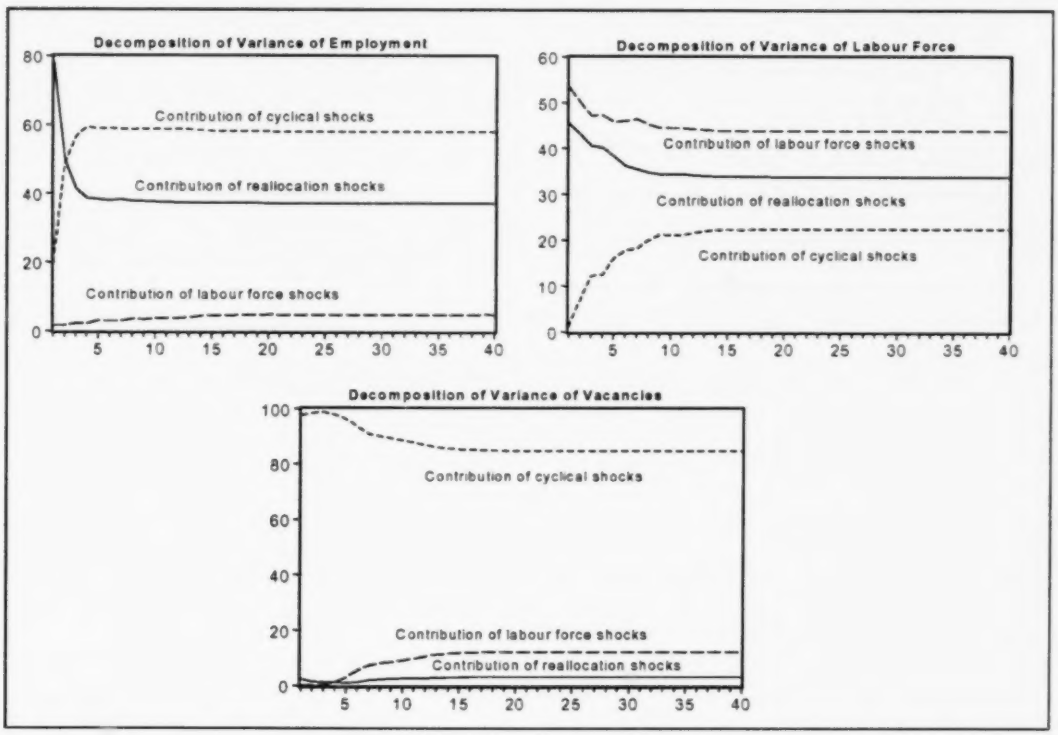
Figure 4 presents the system's impulse responses to the structural shocks. The responses of e and v to the reallocation shocks maintain the theoretical signs for 20 quarters, and this type of shock has permanent effects only on v. One also notes that the aggregate activity shocks and the labour force shocks have permanent effects on the level of all of the variables. The only responses whose sign is not consistent with expectations are those of e and v to labour force shocks.

Figure 4 Impulse Responses to Structural Shocks


Cumulative responses of employment to a reallocation shock
Cumulative responses of employment to a cyclical shock
Cumulative responses of employment to a participation rate shock
Cumulative responses of vacancies to a reallocation shock
Cumulative responses of vacancies to a cyclical shock
Cumulative responses of vacancies to a participation rate shock
Cumulative responses of unemployment to a reallocation shock
Cumulative responses of unemployment to a cyclical shock
Cumulative responses of unemployment to a participation rate shock

The cyclical shocks have a greater impact than the other shocks. However, the relative size of the responses depends on the normalization of the shocks. It is the variance decomposition, presented in Figure 5, that indicates which shock dominates variability. Reallocation shocks contribute to most of the employment fluctuations up to a horizon of 3 quarters. At longer horizons, their contribution is surpassed by that of cyclical shocks, which explain about 60% of employment variability. Labour force shocks contribute to less than 10% of employment fluctuations. The variability of vacancies is totally dominated by cyclical shocks, which explain almost 100% of the variation at short horizons and whose proportion stabilizes at 85% after 15 quarters. Labour force variability is dominated by shocks to the participation rate, although cyclical shocks have less impact than reallocation shocks. Since the variability of V is dominated by cyclical shocks and only the movements of this variable induce a system adjustment dynamic, one must expect cyclical shocks to be the only ones that create a correction mechanism and whose impacts on the unemployment rate are temporary. Given their inability to influence V, the s and f shocks should have a more lasting, if not permanent, impact on the unemployment rate.

Figure 5 Variance Decomposition


Decomposition of Variance of Employment
Contribution of cyclical shocks
Contribution of reallocation shocks
Contribution of labour force shocks
Decomposition of Variance of Labour Force
Contribution of labour force shocks
Contribution of reallocation shocks
Contribution of cyclical shocks
Decomposition of Variance of Vacancies
Contribution of cyclical shocks
Contribution of labour force shocks
Contribution of reallocation shocks

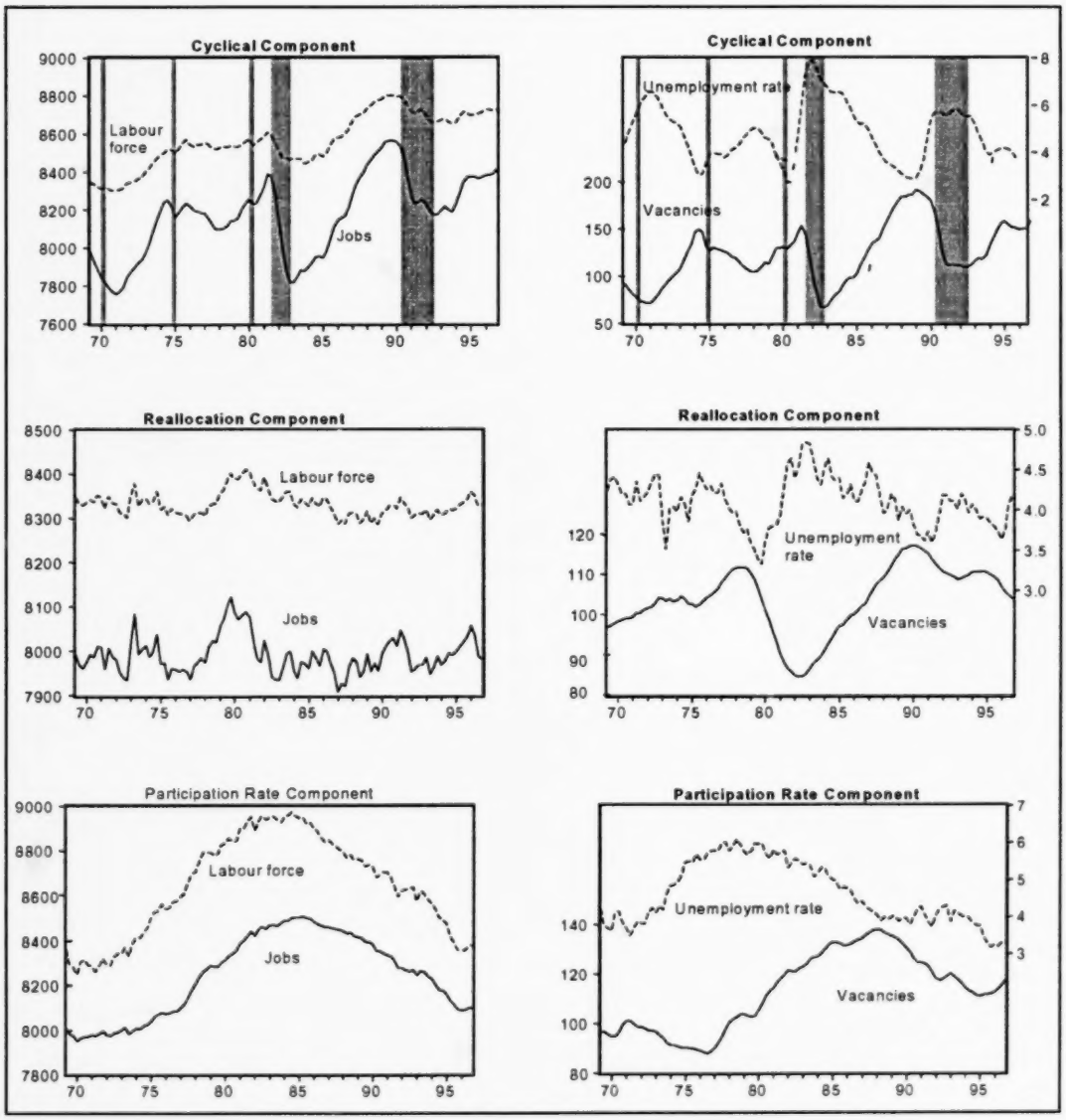
The historical decomposition serves to measure the contribution of these shocks to the observed movements of the variables. Figure 6 shows the impact of the different types of shocks on employment fluctuations, vacancies and unemployment. Since the model is estimated in differences with an error correction term, it decomposes the variation of variables e, v and l into four components, namely those deriving from the three structural shocks and the impact of the deterministic variables. The variations have been cumulated and attached to the initial level of the series. Here we briefly review the principal facts that emerge.

Cyclical shocks largely dominate the short-run variability of the series. They coincide with the dates of business cycles, as shown by the shaded bands identifying periods of recession. Furthermore, one notes that this type of shock has had permanent effects on the levels of E and L, but the effects on the unemployment rate are merely transitory. The model attributes an increase of 5 percentage points in the unemployment rate during the 1981-82 recession and of 3 points during the 1990-92 recession to cyclical shocks, even though the periods of stronger cyclical tensions were just prior to these two recessions, in 1980 and 1989. The contribution of reallocation shocks is weak, both for labour force and employment, such that the strongest impact on unemployment was felt between 1979 and 1982, with an increase in the unemployment rate of about 1.2 percentage units. The weak impact of sectoral reallocation shocks that we observe is consistent with BD's findings in the United States.

The estimated impact on labour force and employment of shocks to the participation rate is the most distinctive element of our findings compared with those of BD. Starting in 1972, the labour force begins to grow at a faster pace, culminating in 1983 with an increase of nearly 600,000 persons. Subsequently this movement begins to reverse, and the labour force slowly declines through the eighties, with the decrease accelerating as of 1993. Employment shows a similar movement but an amplitude that is a little less pronounced, and this series reacts with a slight lag so that the job increase does not start until 1974 and the peak effect is in 1984. Given the time lags between the two series and the elevation of their levels during the seventies, the impact on the unemployment rate is not exactly consistent with those for the participation rate. The rise in the participation rate created a peak effect on the unemployment rate in the third quarter of 1978, with an increase of 2.6 percentage units compared with the level reached in 1971. Such a finding lends credit to the thesis that the participation rate reacted strongly to the 1971 unemployment insurance

reform and to the subsequent adjustments designed to make the program less generous. For example, between 1993:2 and 1996:2, the labour force declined by over 250,000 persons, contributing to a drop in unemployment of nearly 1 percentage unit.

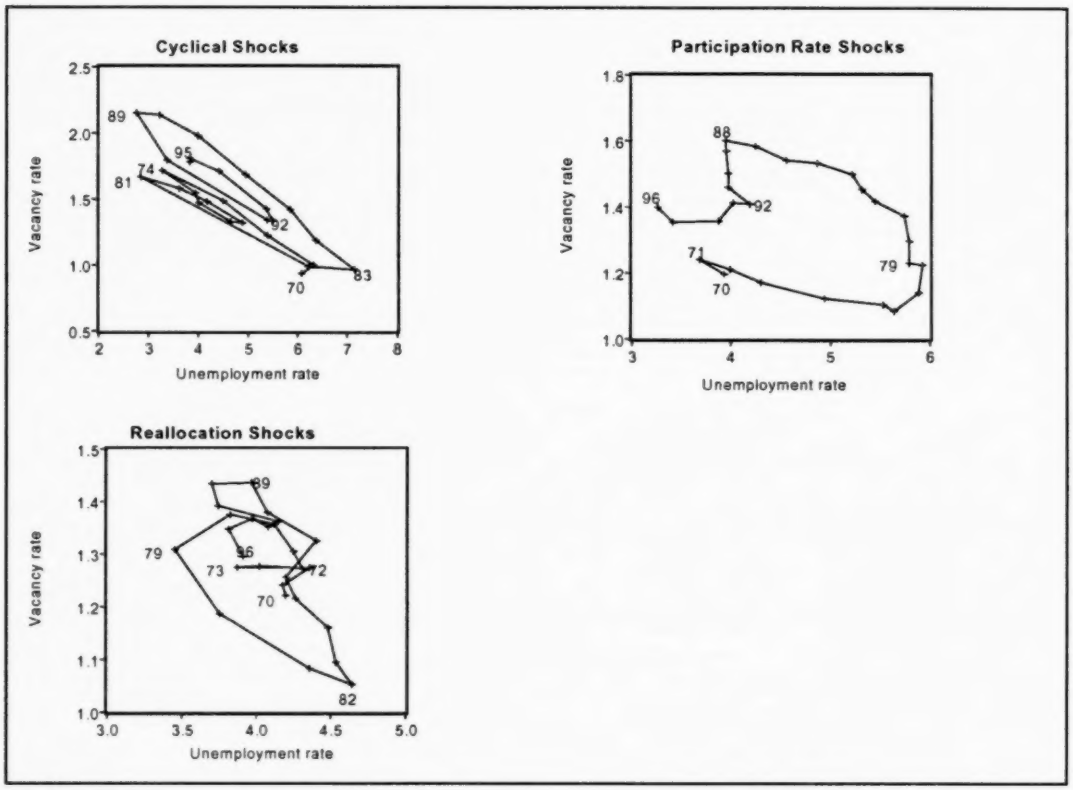
Figure 6 Historical Decomposition



Another important result of the model is the estimated trend contribution. BD had found a trend increase of about 2% between the 1950s and 1980s in the United States. Since the deterministic growth rate of the labour force is higher than that for employment, the model implies a faster trend increase of the unemployment rate, which rose from 4% in 1970 to over 10% in 1996. In view of this strong contribution, it would be interesting to know the reasons for this trend increase. However, the approach we have taken uses correlations at high frequencies to identify shocks, and so does not allow us to shed light to the causes for low-frequency movements of the series. Thus it appears to be difficult to clarify what may have led to such a strong trend increase in the unemployment rate. There might have been a slow but significant deterioration of the degree of agreement between jobs created and jobs destroyed, or a progressive change in the composition of the Canadian labour force.

Figure 7 presents certain results of the historical decomposition in the Beveridge space. Cyclical shocks produce shifts strongly concentrated along a negative curve whose slope shows that an increase of 1% in the unemployment rate, induced by a cyclical shock, reduces the vacancy rate by 0.14%. Reallocation shocks create long circular movements in the Beveridge space whose duration is approximately 15 years. However, the amplitude of the movement on the unemployment rate is relatively weak compared with the cyclical shocks. Finally, labour force shocks produce a large counterclockwise loop lasting about 25 years. Interestingly enough, the final position for 1996 is almost identical to that for 1970.

Figure 7 Structural Shocks in the Beveridge Space

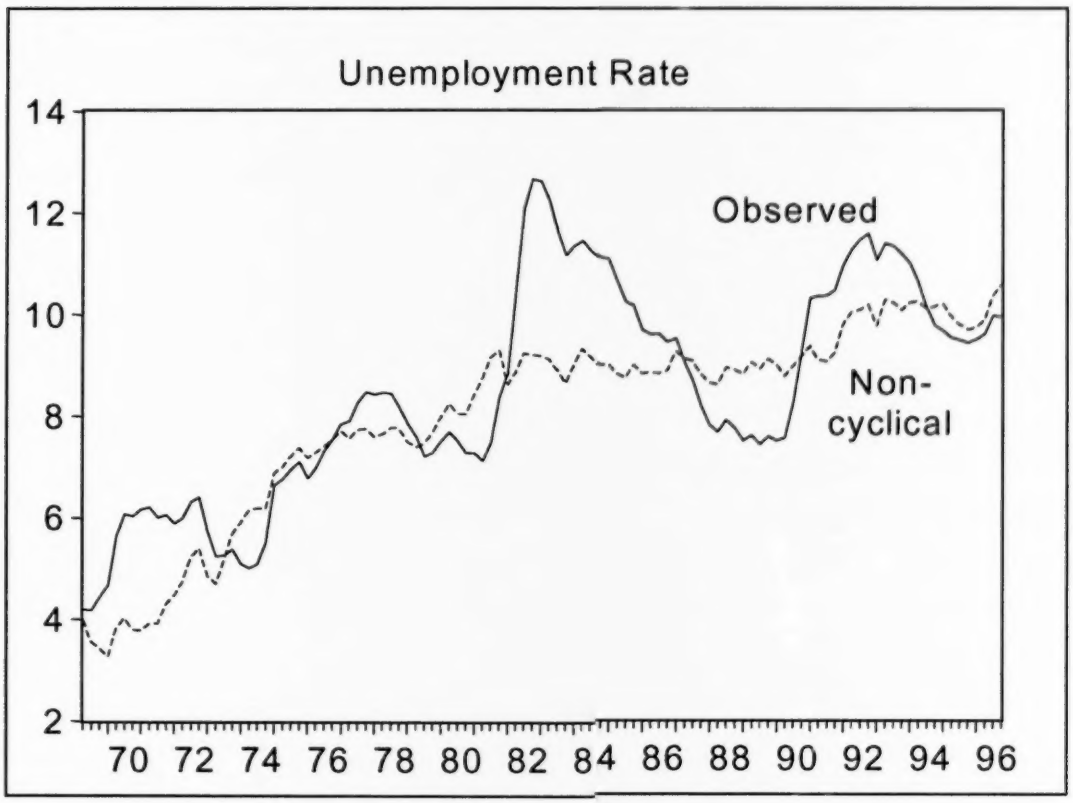


Given the different concepts for defining a natural rate of unemployment, it is not easy to directly relate the model's results to that rate. One well-accepted definition is to calculate an unemployment rate that is free of cyclical fluctuations. This is shown in Figure 8. To obtain this figure, one subtracts the estimated cyclical component from the observed unemployment rate.[24] The model indicates that the non-cyclical unemployment rate rose quickly between 1971 and 1981, from 4% to 9%, and has since increased to about 10%. This is a fairly high estimate compared with those generated by Phillips curves. If this estimate cannot be interpreted as a measurement of the

[24] Remember that, to present the series in level form, we had attached them to the initial level of the series. The cyclical component of the unemployment rate is in fact nearly centred on zero. The non-cyclical unemployment rate is therefore the sum of the deterministic, reallocation and participation rate components.

natural unemployment rate because of the absence of methodological links to the inflation rate, it nonetheless seems consistent with the appearance of inflationary pressures in the late eighties and the relative stability of inflation since 1992.

Figure 8 Non-Cyclical Unemployment Rate



5. Conclusion

The main question dealt with in this paper is to estimate the contribution of participation rate shocks to variations in the Canadian unemployment rate. We conclude that this contribution was substantial in the late seventies and early eighties, with a cumulative effect of over 2.5 percentage points. Although the model is silent on the factors that might have contributed to raising the participation rate in this way, we can identify at least two: the changes made by the 1971 reform of the unemployment insurance system, and the massive influx of baby boomers onto the labour market. The culmination date and general form of the effect of the labour force shocks in fact coincide with the entry of young people into the labour force. However, the current level of unemployment cannot be explained by such shocks, since their effect has progressively dissipated since 1982.

Reallocation shocks play a more minor role in fluctuations of the unemployment rate. Furthermore, their action is slow to emerge and to dissipate. In short, their contribution to short-run variations in unemployment is negligible. As for cyclical shocks, their action dominates the short-run variability of the unemployment rate, and they are responsible for its major fluctuations. The explanation for the high level of unemployment in 1996 compared with 1970 remains obscure, however, since the model attributes this to deterministic factors which led the labour force to grow faster than employment. Our identification hypotheses do not allow us to further analyse this type of movement.

Our work has also revealed that labour market imbalance is directly related to the unemployment rate, and that only the creation of vacancies brings about an imbalance correction dynamic. The reduced-form model in fact demonstrates that vacancies alone react to the imbalance, and only their action is capable of modifying the time path of employment and labour force. Without this corrective action, the dynamic behaviour of the unemployment rate would conform to the hysteresis hypothesis. If we tie this to the fact that the variability of the shocks to V is completely dominated by cyclical shocks, we have a demonstration that only cyclical shocks have a temporary impact on the unemployment rate. Reallocation and participation rate shocks, for their

part, have an impact on unemployment that is permanent. This work suggests that the time has come to stop neglecting vacancies in the analysis of unemployment.

Our approach also makes intensive use of the Help-Wanted Index as a proxy for a genuine vacancy index. Since the last direct estimates of vacancies date back almost 20 years, it is possible that our estimate of the vacancy rate is currently biased. We believe that this affects our conclusions only slightly, since our model permits independent trend movements in each series. If the HWI poorly captures the inherent vacancy trend, as Abraham's study (1987) suggests, this should be translated chiefly by a different constant in the vacancy equation so that the non-deterministic components would have to remain similar. Nonetheless, in view of the role we assign them in the correction of disequilibrium, it is still important for a survey on vacancies to be carried out.

In closing, given that we have substantial regional disparities in Canada, it would be interesting to redo this analysis at a disaggregated level. It could be repeated with the five major Canadian regions (the Atlantic provinces, Quebec, Ontario, the Prairies and British Columbia) and with certain age groups at the national level. Among other things, one might well think that labour force shocks have a more significant impact in the Atlantic provinces than in Ontario, or that their impact is greater among those 15 to 24 years old than those 25 to 54 years of age. This is an interesting avenue for future research which would help us to better understand the dynamics of unemployment fluctuations in Canada.

Appendix 1

Consolidation of the Help-Wanted Index

Although the two Help-Wanted Indices (HWI) measures are related, they are also conceptually distinct. One result of this is that the new series behaves more smoothly than the old one at high frequencies. We preferred to consolidate the series by making a reverse forecast from the new series rather than a forecast from the old one since 1981, for two reasons. First, the relevance of the HWI as a vacancy rate proxy has been demonstrated by Abraham (1987) using the number of Help-Wanted ads. Hence it seemed to us preferable to use the same concept. Furthermore, by doing the forecast backward, new observations can be automatically added to the consolidated series instead of continually adding forecasts in place of new observations.

The method adopted for effecting the forecast is to estimate a link capable of producing prediction errors during the period common to the two series which are not serially correlated. To do this, we estimated on monthly data the logarithm of the new index LHWIN on the logarithm of the old index LHWIO. Since this simple regression produced autocorrelated errors, we progressively added lead values, and at each stage the equation was subjected to LM tests at 4 and 8 lags to detect the presence of autocorrelation of the residuals. With 3 lags, no serial correlation was detectable. Analysis by recursive least squares was then used to test temporal stability, and only late 1982 seems subject to stability problems. Since this problem could not be remedied by extending the lag structure, we adopted the equation with 3 lags. This equation is LHWIN = 0.108 + 1.135*LHWIN(+1) - 0.154*LHWIN(+2) - 0.062*LHWIN(+3) + 0.197*LHWIO - 0.0117*LHWIO(+1) - 0.047*LHWIO(+2) - 0.073*LHWIO(+3). We then did a reverse flow forecast, and the series obtained was returned to level form by an inverse logarithmic transformation. After the consolidation, variability at high frequencies becomes almost identical before and after 1981 since the standard deviation of the change in the series variation rate is 3.69% before 1981 and 4.05% from 1981 to 1996.

Appendix 2

Results of Estimation of the Reduced-Form Model

Cointegration relation normalized relative to LE :

LE(-1) - 0.009*LV(-1) - 0.839*LL(-1)

Dependent Variable DLE - Estimation by Least Squares
Quarterly Data From 1967:04 To 1996:04
Usable Observations 117
Degrees of Freedom 97
Centered R**2 0.674016
R Bar **2 0.610164
Uncentered R**2 0.797841
T x R**2 93.347
Mean of Dependent Variable 0.0050353950
Std Error of Dependent Variable 0.0064616166
Standard Error of Estimate 0.0040344306
Sum of Squared Residuals 0.0015788332

Variable	Coeff	Std Error	T-Stat	Signif
1. DLE{1}	-0.151380768	0.183109318	-0.82672	0.41042286
2. DLE{2}	0.321208147	0.184651061	-1.73954	0.08511165
3. DLE{3}	-0.099186495	0.185119895	-0.53580	0.59332602
4. DLE{4}	-0.082591353	0.177334811	-0.46574	0.64244795
5. DLE{5}	-0.146810212	0.171715129	-0.85496	0.39467799
6. DLE{6}	0.067636246	0.155697332	0.43441	0.66495663
7. DLV{1}	0.094789391	0.012979621	7.30294	0.00000000
8. DLV{2}	0.012335541	0.017614206	0.70032	0.48540524
9. DLV{3}	0.004394610	0.017830282	0.24647	0.80584010
10. DLV{4}	-0.023243135	0.017847659	-1.30231	0.19589516
11. DLV{5}	0.032975360	0.017985996	1.83339	0.06981122
12. DLV{6}	-0.003018210	0.015959582	-0.18912	0.85039715
13. DLL{1}	0.178798720	0.201064929	0.88926	0.37606416
14. DLL{2}	0.188107195	0.19617978	0.95885	0.34001769
15. DLL{3}	0.147575788	0.190698589	0.77387	0.44089009
16. DLL{4}	0.024579586	0.187686408	0.13096	0.89607745
17. DLL{5}	0.374198847	0.186962857	2.00146	0.04813615
18. DLL{6}	0.233230508	0.179570662	1.29882	0.19708376
19. Constant	-0.058234086	0.070269226	-0.82873	0.40929282
20. RELCO{1}	0.042109609	0.048496996	0.86829	0.38737708

Appendix 2 (Continued)

Dependent Variable DLV - Estimation by Least Squares
Quarterly Data From 1967:04 To 1996:04
Usable Observations 117
Degrees of Freedom 97
Centered R**2 0.715406
R Bar **2 0.659660
Uncentered R**2 0.715408
T x R**2 83.703
Mean of Dependent Variable 0.0001708613
Std Error of Dependent Variable 0.0611124041
Standard Error of Estimate 0.0356521471
Sum of Squared Residuals 0.1232943327

Variable	Coeff	Std Error	T-Stat	Signif
1. DLE{1}	0.271668499	1.618131770	0.16789	0.86701919
2. DLE{2}	0.868952157	1.631756106	0.53253	0.59557970
3. DLE{3}	0.797104112	1.635899181	0.48726	0.62717523
4. DLE{4}	1.521836643	1.567102619	0.97111	0.33390644
5. DLE{5}	0.189689686	1.517441644	0.12501	0.90077731
6. DLE{6}	-0.104858207	1.375892833	-0.07621	0.93940818
7. DLV{1}	0.857600771	0.114700539	7.47687	0.00000000
8. DLV{2}	-0.176617263	0.155656229	-1.13466	0.25931299
9. DLV{3}	-0.096212454	0.157565691	-0.61062	0.54288058
10. DLV{4}	-0.011222655	0.157719248	-0.07116	0.94342018
11. DLV{5}	-0.097412462	0.158941724	-0.61288	0.54138891
12. DLV{6}	0.085034293	0.141034367	0.60293	0.54796029
13. DLL{1}	0.909991331	1.776804986	0.51215	0.60970991
14. DLL{2}	-1.507681521	1.733635114	-0.86966	0.38663052
15. DLL{3}	-1.951559140	1.685197940	-1.15806	0.24968325
16. DLL{4}	-1.854961102	1.658579381	1.11840	0.26615680
17. DLL{5}	-1.320282003	1.652185376	-0.79911	0.42617759
18. DLL{6}	-1.059207417	1.586860768	-0.66749	0.50604664
19. Constant	1.602086213	0.620967122	2.57999	0.01138018
20. RELCO{1}	-1.100633113	0.428566560	-2.56817	0.01174919

Appendix 2 (continued)

Dependent Variable DLL - Estimation by Least Squares
Quarterly Data From 1967:04 To 1996:04
Usable Observations 117
Degrees of Freedom 97
Centered R**2 0.458614
R Bar **2 0.352570
Uncentered R**2 0.794674
T x R**2 92.977
Mean of Dependent Variable 0.0056285652
Std Error of Dependent Variable 0.0044184978
Standard Error of Estimate 0.0035552582
Sum of Squared Residuals 0.0012260665

Variable	Coeff	Std Error	T-Stat	Signif
1. DLE{1}	-0.129023990	0.161361286	-0.79960	0.42589813
2. DLE{2}	-0.030676959	0.162719915	-0.18853	0.85085808
3. DLE{3}	-0.074466319	0.163133066	-0.45648	0.64906825
4. DLE{4}	0.072309719	0.156272622	0.46272	0.64460487
5. DLE{5}	-0.185899673	0.151320393	-1.22852	0.22222530
6. DLE{6}	0.010903716	0.137205042	0.07947	0.93682233
7. DLV{1}	0.030287880	0.011438022	2.64800	0.00945154
8. DLV{2}	0.009937192	0.015522153	0.64019	0.52355572
9. DLV{3}	-0.010543964	0.015712566	-0.67105	0.50378163
10. DLV{4}	-0.025426265	0.015727879	-1.61664	0.10920450
11. DLV{5}	0.043214597	0.015849785	2.72651	0.00759670
12. DLV{6}	-0.005945980	0.014064050	-0.42278	0.67339256
13. DLL{1}	0.127120016	0.177184296	0.7174	0.47482367
14. DLL{2}	0.008893098	0.172879365	0.05144	0.95907985
15. DLL{3}	0.130174911	0.168049174	0.77462	0.44044604
16. DLL{4}	-0.107794905	0.165394752	-0.65174	0.51610812
17. DLL{5}	0.441960896	0.164757137	2.68250	0.00859104
18. DLL{6}	0.247437286	0.158242919	1.56365	0.12115542
19. Constant	0.030386278	0.061923297	0.49071	0.62474103
20. RELCO{1}	-0.019451411	0.042736972	-0.45514	0.65002383

Bibliography

Abraham, K.G. "Help-Wanted Advertising, Job Vacancies, and Unemployment." *Brookings Papers on Economic Activity*, 1:1987: 207-248.

Abraham, K.G. and L.F. Katz. "Cyclical Unemployment: Sectoral Shifts or Aggregate Disturbances?" *Journal of Political Economy*, 94, 3 (1986): 507-522.

Blanchard, O.J. "Les Courbes de Beveridge et de Phillips Comme Outils d'Analyse du Chômage." *L'Actualité Économique, Revue d'Analyse Économique*, 65, 3 (September 1989): 396-422.

Blanchard, O.J. and P. Diamond. "The Beveridge Curve." *Brookings Papers on Economic Activity*, 1:1989: 1-76.

Burns, Andrew. "Unemployment in Canada: Frictional, Structural and Cyclical Aspects." Working Paper No. 1. Economic Council of Canada, 1990.

Card, David, and W. Craig Riddell. "A Comparative Analysis of Unemployment in Canada and the United States." Working Paper 297. Industrial Relations Section, Princeton University, 1992.

Coe, David T. "Structural Determinants of the Natural Rate of Unemployment in Canada." *IMF Staff Papers*, 37 (1990): 94-115.

Cross, P. "Alternative measures of business cycles in Canada: 1947-1992." *Canadian Economic Observer*, Statistics Canada, Cat. 11-010-XPB, February 1996, 3.1-3.40.

Davis, S. and J. Haltiwanger. "Gross Job Creation, Gross Job Destruction, and Employment Reallocation." National Bureau of Economic Research, Working Paper No. 3728, 1991.

Doornik, J. A. and H. Hansen. "An Omnibus Test for Univariate and Multivariate Normality." Working Paper. Nuffield College, Oxford, 1994.

Fillion, J. F. and A. Léonard. "La courbe de Phillips au Canada: un examen de quelques hypothèses." Discussion Paper 97-3, Bank of Canada, 1997.

Fortin, M. and A. Araar. "Sectoral Shifts, Stock Market Dispersion and Unemployment in Canada." *Applied Economics*, 29 (June 1997).

Fortin, M. and Fortin, P. "Les déterminants du taux d'activité au Canada: une approche par groupes démographiques et par régions." Paper presented to the congress of the SCSE, École des hautes études commerciales, Montreal, May 14, 1997.

Fortin, P. "The Great Canadian Slump." *Canadian Journal of Economics*, 29, 4 (1996): 761-787.

Fortin, P. "Unemployment Insurance Meets the Classical Labor Supply." *Economics Letters*, 14 (1984): 275-281.

Hansen, H. and K. Juselius. *Cats in Rats: Cointegration Analysis of Time Series*. Evanston, 1995.

Johansen, S. "Statistical Analysis of Cointegration Vectors." *Journal of Economic Dynamics and Control*, 12 (1988): 231-254.

Jones, Stephen R.G. and W. C. Riddell. "The Measurement of Unemployment: An Empirical Approach." Working Paper No. 93-48. Department of Economics, University of British Columbia, 1993.

King, R., C. Plosser, J. Stock and M. Watson. "Stochastic Trends and Economic Fluctuations." *American Economic Review*, 81 (1991): 819-840.

Lilien, D.M. "Sectoral Shifts and Cyclical Unemployment." *Journal of Political Economy*, 90, 4 (1982): 777-793.

McCallum, J. "Les taux de chômage canadien et américain dans les années 1980: un test de trois hypothèses." *L'Actualité économique, Revue d'analyse économique*, 64, 4: 494-508.

Milbourne, R.D., D.D. Purvis and W.D. Scoones. "Unemployment Insurance and Unemployment Dynamics." *Canadian Journal of Economics*, XXIV, 4 (November 1991): 804-826.

Moorthy, V. "Unemployment in Canada and the United States: The Role of Unemployment Insurance Benefits." *Federal Reserve Bank of New York Quarterly Review* (Winter 89-90), p. 48-61.

Phillips, P.C.B. "Impulse Response and Forecast Error Variance Asymptotics in Nonstationary VAR's." Discussion Paper No. 1102, Cowles Foundation for Research in Economics, Yale University, June 1995.

Pissarides, C. A. *Equilibrium Unemployment Theory*. Oxford: Basil Blackwell, 1990.

Samson, L. "A Study of the Impact of Sectoral Shifts on Aggregate Unemployment in Canada." *Canadian Journal of Economics*, XVIII, 3 (August 1985): 518-530.

Samson, L. "The Beveridge Curve and Regional Disparities in Canada." *Applied Economics*, 26 (1994): 937-947.

Setterfield, M.A., D.V. Gordon and L. Osberg. "Searching for a Will o' the Wisp: An Empirical Study of the NAIRU in Canada." *European Economic Review*, 36 (1992): 119-136.

Storer, P. "Persistent Unemployment and Sectoral Shocks in the Wake of the 1982 Recession." Cahier de recherche du CREFE, Université du Québec à Montréal, May 1992.

Zagorsky, Jay L. "Job Vacancies in the United States and Canada." *Journal of Economic and Social Measurement*, 19 (1993): 305-319.